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Returns to Education in the Economic Transition: A Systematic Assessment Using Comparable Data[1]

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Abstract

This paper examines the assertion that returns to schooling increase as an economy transitions to a market environment. This claim has been difficult to assess as the existing empirical evidence covers only a few countries over short time periods. A number of studies find that returns to education increased from the "pre-transition" period to the "early transition" period; it is not clear what has happened to the skills premium through the late 1990s, or the period thereafter. We use data that are comparable across countries and over time to estimate returns to schooling in eight transition economies (Bulgaria, Czech Republic, Hungary, Latvia, Poland, Russia, Slovak Republic and Slovenia) from the early transition period up to 2002; in the case of Hungary, we capture the transition process more fully, beginning in the late 1980s. Compared to the existing literature, we implement a more systematic analysis and perform more comprehensive robustness checks on the estimated returns, although at best we offer only an incomplete solution to the problem of endogeneity. We find that the evidence of a rising trend in returns to schooling over the transition period is generally weak, except in Hungary and Russia where there have been sustained and substantial increases in returns to schooling. On average, the estimated returns in our sample are comparable to advanced economy averages. There are, however, significant differences in returns across countries and these differentials have remained roughly constant over the last 15 years. We speculate on the likely institutional and structural factors underpinning these results, including incomplete transition and significant heterogeneity and offsetting developments in returns to schooling within countries.

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1. Introduction

Under central planning, returns to schooling were typically low and differences in educational attainment had limited impact on individual variations in earnings. Labor market liberalization, decentralized wage setting and the broader transition to a market economy are assumed to lead to an increase in wage differentials and earnings inequality. A number of factors may however offset the expected increases in the returns to schooling over the transition period. For example, skills and experience acquired under the communist system may be less marketable in the new market environment compared to human capital acquired during the transition.[2] This hypothesis implies that the transition to a market economy can be associated with some offsetting decline in returns to schooling.

Other developments over the transition period may serve to either promote or dampen the increase in the skill wage premium. For example, falling government expenditures on education, under a broad fiscal adjustment strategy, may lower the quality of education (Campos and Jolliffe 2004). Supply factors also play an important role and economies in transition typically have a relatively highly educated labor force. Furthermore, changes in labor force participation may also either exacerbate or offset rising wage inequality.[3] Finally, the speed and uniformity of transformation may matter. As documented in the broader literature on transition economies, the pace of privatization and enterprise restructuring, among other structural reforms, has differed widely across countries. Where the new private sector remains a small share of total economic activity, wage-setting mechanisms inherited from the socialist era may persist. Taking into account all these possible interactions, it is not obvious that the transition to a market economy necessarily leads to sustained increases in returns to schooling.

Previous attempts to evaluate the issue empirically show mixed results and inconclusive evidence. A number of studies find that returns to education increased from the "pre-transition" period to the "early transition" period; it is not clear what has happened to the skills premium through the late 1990s, or the period thereafter. We believe that the existing literature is constrained by at least three major weaknesses (i) the time span considered is too short; (ii) the data used to assess the evolution over time both within and across countries are from datasets of different quality and sample design; (iii) robustness of results is not sufficiently investigated. This paper attempts to resolve these issues.

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[2] See Munich and others (2004) for a discussion of this hypothesis.

[3] In Poland, Newell and Socha (2005) find that the share of employees with relatively lower educational attainment has fallen over time; they show that estimated increases in returns to education are not robust to shifting patterns of participation.

Using data drawn from the *International Social Survey Programme* (ISSP), we cover eight transition economies (Bulgaria, Czech Republic, Hungary, Latvia, Poland, Russia, Slovak Republic, and Slovenia) over a relatively long time period: from the early transition period to 2002, although we are unable to capture the transition process fully except in the case of Hungary, where data are available beginning in the late 1980s. This allows us a rare, medium-term perspective necessary to study structural labor market adjustments over time as the homogeneity of the data is unprecedented. The ISSP data allow us to assess both cross-country heterogeneity and structural breaks in returns to schooling.[4]

We also assess the robustness of our results in a fairly comprehensive way by testing their sensitivity to different specifications and to the estimation techniques applied. These robustness checks are crucial as they show that results may change quite radically. For example, in the basic specification with OLS—a standard Mincerian regression—we find an increasing trend in returns to education among countries for which we have a longer time span (Hungary and Russia). However, a more complete specification yields a much weaker trend. Cross-country rankings, in contrast, are generally robust to the specification and to gender/cohort heterogeneity.

The fundamental issue of endogeneity bias in returns to schooling can only be partially assessed with the data at hand. We use additional information on family background, available for about one or two years for each country, to perform some control function estimation on returns to college education. What we find is consistent with the recent literature: OLS estimates are downward biased.[5] Given the limited information available, we are, however, unable to fully evaluate whether the downward bias is crucial to explaining trends, or lack thereof. We can nevertheless confirm that cross-country rankings of returns to schooling are roughly consistent across estimation techniques.

Finally, we perform cross-country regressions to investigate the correlation of returns to schooling with measures of reform progress, conditioning on country fixed-effects, time trend and macroeconomic controls. The positive correlation usually found between reform speed and returns to schooling is not very robust.

The remainder of the paper is organized as follows: Section 2 reviews the relevant literature. Section 3 describes the data source. The econometric specification is presented in Section 4. Section 5 reports on and discusses the returns to schooling drawn from the OLS results. Section 6 presents the results from the control function estimates of returns to

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[4] Data from ISSP lend themselves to meaningful cross-country comparisons not just for economies in transition but for other country groups as well. For example, Blau and Kahn (2000) use ISSP data to compare gender earnings differentials across selected OECD countries.

[5] See for example the surveys in Card (2001) and Card (1999).

College education and Section 7 from the cross-country regressions. Section 8 summarizes the main results and policy implications of the paper.

2. Brief Review of Related Literature

The empirical literature on returns to schooling in transition countries has tended to cover only a few countries over relatively short time periods. Over the last decade, for example, a large literature on returns to education or schooling has emerged for the Czech Republic, Hungary, Poland, and Russia, but relatively less is known about other countries such as Bulgaria, Latvia, the Slovak Republic and Slovenia and, more generally, very little is known about many other transition economies (see Psacharopolous and Patrinos 2004 for a recent survey of the literature).

In addition, few country-specific empirical studies investigate patterns of returns to education or schooling after the early transition phase or after the early or mid-1990s (see Appendix Table 1 for summary information on selected studies of the countries in our sample). There are some exceptions: Munich and others (2004) use Czech data for 1996 and 2002, with retrospective data for 1989. Kertesi and Köllõ (2001) use firm-level surveys in Hungary over the 1986 to 1999 period. Newell and Socha (2005) use Poland's *Labor Force Survey* (LFS) data covering the 1994 to 2002 period. The World Bank (2003) uses data from the RLMS to estimate the returns to schooling in Russia over the 1992 to 2000 period.

Cross-country studies using comparable data and methodology to evaluate the levels and trends in returns to schooling in transition economies are also rare. Fleisher and others (2005) use metadata of returns to schooling between 1975 and 2002 collected from 33 studies of 10 transition economies, including China, to assess changes in the returns to schooling over the transition period.

Both country-specific and cross-country studies find that returns to education increased from the "pre-transition" period to the "early transition" period.[6] The meta-study by Fleisher and others (2004) also suggests that the sharpest increases in returns to education took place during the early transition (this period varies from country to country but is roughly around the early 1990s), although there is evidence that returns to schooling continued rising after this initial period.

In general, however, it is less clear in the literature what has happened to the skills premium in transition economies through the late 1990s, or the period thereafter (through 2002 or 2003). Despite the early literature documenting the rising returns to education, a number of studies suggest that returns may have remained stable after the early transition period. In Poland, for example, Rutkowski (2001) finds that the increase in returns to

[6] See also Rutkowski (2001) who reviews some of the related country-specific literature. See also Appendix Table 1.

education occurred mostly in the early transition period. From 1993 through the late 1990s, the education premium remained largely unchanged. Similarly, Vodopivec (2004) suggests that in Slovenia, most of the striking change in returns to education took place prior to 1993. In the Czech Republic, Munich and others (2004) report substantial increases in returns to education for women in the Czech Republic between 1989 and 1996; between 1996 and 2002, however, no further change is observable. Meanwhile, Kertesi and Köllõ (2001) found an aggregate increase in returns to education in Hungary between 1989 and 1992; however, the productivity and wages of older workers apparently did not grow any further after 1992. In Russia, Cheidvasser and Benitez-Silva (2000) are unable to find increases in returns to education between 1992 and 1999, in contrast to other studies who have found rising returns over this same period (see for example World Bank 2003).

With respect to cross-country rankings, Fleisher and others (2004) suggest that cross-country differences in returns to schooling are persistent. However, other than this meta-study, it is difficult to draw conclusions about cross-country differences in trends and levels of returns to schooling owing to many technical issues related to comparability.

In general, existing studies suffer from a number of shortcomings, with respect to the ability to evaluate returns to schooling over the medium-term. First, most studies cover short time periods and few studies cover the transition period through the late 1990s. Second, cross-country comparisons are from datasets of different quality and sample design. This makes it difficult to draw conclusions about differences in returns to education across countries, as the conclusions may be contaminated by cross-country data that are not, strictly speaking, comparable. Fleisher and others (2005) attempt to control for differences in sample design, but their metadata are nonetheless drawn from various studies using different datasets and specifications. Third, in some cases, the same country is examined over time using datasets of different quality and sample design. For example, a study of the Czech and Slovak Republics (Chase 1998) utilizes data for 1984 and 1993 drawn from two distinct data collection efforts that implement different sampling frames, with hours of work and earnings measured differently. The author carefully attempts to render the two surveys as comparable as possible but the results are nonetheless likely to be subject to some degree of error. Pre-transition information based on retrospective data may also be subject to recall measurement error. Finally, in many studies, the robustness of results is not sufficiently investigated as OLS estimation techniques are employed without accounting for possible biases.

The use of comparable, cross-national individual-level surveys, as described in the next section, is therefore particularly promising to address some of these shortcomings.

3. Data

For our analysis we use data from the *International Social Survey Program* (ISSP). The ISSP is an ongoing annual program involving collaborative, international survey data collection efforts. The program has been conducted annually since 1985, covering selected topics in the social sciences. The surveys initially covered a narrow sample of industrial countries but, over time, have expanded to include more transition economies as well as developing countries. The ISSP currently consists of 32 national cross-sectional surveys. The critical characteristic that makes them appropriate for our purposes is that ISSP surveys are based on a common sampling and methodological framework and are thus comparable both between and within countries over time. The sample stratification generates nationally representative samples and provides individual-level information on demographic and socio-economic characteristics and personal views on selected social topics for at least 1,000 respondents per country per year. After excluding self-employed workers, retirees, and students, our net sample size is about 500 individuals (age 18 to 65) per country per year.

We create a database covering over 70 individual-level survey datasets for 8 transition economies (Bulgaria, Czech Republic, Hungary, Latvia, Poland, Russia, Slovak Republic, and Slovenia) over the 1986 to 2002 period, although Hungary alone has enough relevant information before the early 1990s. In particular we cover the following periods for each country: Hungary (1986-2002), Poland (1991-2002), Russia (1991-2002), Bulgaria (1992-2002), Czech Republic (1993-2002), Slovakia (1995, 1998), Slovenia (1991-1992), and Latvia (1996-2002). The Data Appendix describes the database more fully, the specific survey-years we included and why other years were excluded, the definition and comparability of key variables, and the sample restrictions we imposed.

How appropriate are ISSP data to examine issues in labor economics? A number of empirical papers in labor economics have utilized ISSP data to look at the wage curve (Blanchflower and Oswald 1994), the union wage premium (Blanchflower and Bryson 1997), working hours (Bell and Freeman 1999), gender earnings gap in the OECD (Blau and Khan 1992, 1995, and 2000), returns to education in the U.S. (Card 1999), among others. Two studies have utilized ISSP data to conduct a cross-country study of returns to education: to the best of our knowledge, Lorenz and Wagner (1993) are the first to estimate the returns to education using ISSP data. They focus on the second half of the 1980s using a sample of eight (mostly OECD) countries, including Hungary. Trostel, Walker, and Woolley (2002) estimate returns to education for a sample of 28 countries from 1985 through 1995, including the countries in our sample. Their study however, only reports average returns for the 1985 to 1995 period and the subset of results that briefly analyzes the trend rate of returns includes only 2 of the 8 countries we investigate (Russia and Poland).

We consider the ISSP data particularly appropriate for a cross-country study of returns to education in transition economies, for the following reasons:

First, as previously discussed, few cross-country studies focusing on transition economies exist. These studies are typically based on datasets that are not necessarily comparable, even within countries over time.

Second, to understand the evolution of labor markets in transition economies, data collected using the same survey instrument over the transition period bolster our confidence in the credibility of measured changes. As previously discussed, comparisons over time have often come from data collected from different data sources and one is uncertain whether the documented changes reflect actual changes in the population, or are a function of methodological breaks and revisions in the survey instruments themselves (see Campos and Jolliffe 2003 for a discussion). Even fewer studies cover the transition period through the late 1990s up to early 2000s. While there is some consensus that returns to education increased significantly from the pre-transition period to early transition, the pattern thereafter has not been sufficiently investigated.

Third, previous studies use either hourly or monthly earnings as the dependent variable in the earnings regressions, with strong technical assumptions about the exogeneity of hours of work. ISSP data provide both measures and thus allow for robustness tests to the choice of dependent variable.

Fourth, the use of ISSP data to estimate the returns to schooling in Bulgaria, Latvia and Slovenia represents a contribution to the growing, but still limited, literature on these countries. In general, the literature on returns to schooling in transition economies has covered mostly Central and Eastern European countries and Russia, and this paper is not an exception. Nonetheless, to the best of our knowledge, this study is the first to quantify the returns to years of schooling in Latvia, and Slovenia.[7] With the exception of the work of Rutkowski (1999) and Jones and Simon (2004), this study is also the only other attempt to estimate the returns to schooling in Bulgaria.

Finally, we consider our measure of the number of years of schooling to be more accurate than the one adopted in many other studies (e.g., Flanagan 1998, Vernon 2002, Brainerd 1998), since we use the number reported by the respondents themselves, rather than an imputed value derived from the reported school attainment (see Munich et al 1999).

[7] Vodopivec (2004) looks at returns to educational attainment.

4. Econometric Specification

We estimate a standard earnings regression of the form:

$$\log w_{ijt} = x'_{ijt}\,\beta_{jt} + \gamma_{jt} s_{ijt} + \varepsilon_{ijt} \qquad (1)$$

where w represents monthly earnings, s is years of schooling completed and x' is a vector of additional controls. Individuals are denoted by $i=1,..,N_{jt}$; countries by $j=1,...,M_t$; and year by $t=1,...,T$. The parameter γ is the coefficient of interest and it is often interpreted as the "average rate of return to schooling" or the percentage change in wages to an additional year of schooling. This structural interpretation is motivated by a human capital model over the life cycle and it is the basis for many of the policy implications about the optimal level of schooling.[8] As extensively pointed out in Heckman, Lochner and Todd (2005), this interpretation is valid only under quite strict assumptions. Moreover, many models imply that the years of schooling variable is endogenous and therefore the simple OLS estimator used in the next section may be inconsistent.[9] We provide a more detailed discussion of these issues in section 6; for the time being, we focus on estimating γ over time and for each country. The main motivation of the exercise is to provide a benchmark: OLS estimates of the coefficient on years of schooling drawn from earnings regressions are one of the most popular and comparable measures of returns to skills used in the literature and a relevant variable for policy makers in considering government interventions and reforms.

The first set of estimates is reported in Table 1. For ease of cross-country comparisons and comparisons across specifications, see also Figure 1. It is obtained by estimating equation (1) for each country and year available: in other words we estimate conditioning on (j,t). We propose two specification: a basic specification where x' is defined as years of potential experience (linear and squared), a dummy for male and a constant; and a richer specification that adds dummies for living in urban areas and being married, controls for current job (dummies for occupation[10], public employee, working full-time, membership in a trade union), controls for current family (number of members, dummy for spouse working full-time). We label the first specification as "Basic" and the

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[8] The classic reference for this specification is the book by Mincer (1974) where the additional controls were simply years of work experience (linear and squared) and a constant. A huge amount of regressions based on variation of this specification have been estimated, see Psacharopoulos and Patrinos (2004) for a recent survey.

[9] The first systematic assessment in the literature is Griliches (1977), a recent survey and updated interpretation is Card (2001).

[10] The occupational dummies follow the classification of ISCO occupational categories into five dummy variables as implemented in previous studies using ISSP (e.g., O'Rourke and Sinnott 2002): elementary occupations, plant and machine operators, assemblers, etc., technicians and associate professionals, professionals, and legislators.

second as "Richer." We also add a specification labeled "Basic-Balanced" or "Basic on Balanced Sample." The "Basic-Balanced" specification is a regression run using the basic specification on the estimation sample of the richer specification. Due to missing values, the estimation sample of the richer specification is always smaller than the estimation sample of the basic specification and this estimate is used as a benchmark to judge if differences in returns are due to either sample or specification differences.[11]

The second set of estimates, reported in Table 2, specifically looks at differences in returns over time by testing for structural breaks on time sub samples. For each country we pool years together in three periods: "Pre-transition" (1986-1990), "Early transition" (1991-1996) and "Late transition" (1997-2002). The specifications are the same as those reported previously but now constraining the estimated coefficients to be the same for the same country over the three sub-samples except for the constant that remains year-specific.

The third set of estimates, reported in Table 3, evaluates differences over time directly estimating a time trend. In terms of equation (1), all years are pooled together for each country and the coefficients of the two specifications are constrained to be the same over time except for the constant. A linear and quadratic trend is added to the specification and interacted with years of schooling.

Table 4 reports estimates obtained by quantile regression. The OLS regression is based on the mean of the conditional earnings distribution. This approach assumes that possible differences in terms of the impact of the exogenous variables along the conditional distribution are unimportant. This, however, is an open empirical question. Following Buchinsky (1994), we use quantile regressions to estimate earnings at different points of the conditional earnings distribution.[12] These regressions should yield information on the heterogeneity in returns to education, if any, or whether the returns to schooling vary across conditional quantiles.[13] We can think of conditional quantiles as

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[11] For convenience, in all tables reporting estimation results we provide point estimates and standard errors only for the parameters of interest (usually the returns to schooling.) The complete set of results is available in a web appendix at http://www9.georgetown.edu/faculty/lf74/. Appendix Table 2 reports the descriptive statistics by specification.

[12] Martins and Pereira (2003) is a recent contribution using quantile regression analysis to perform cross-country comparisons in returns to schooling. They identify "skilled" and "unskilled" workers, conditional on the schooling and work experience of these workers and they then compare the education premia for these two groups of workers examining how schooling may be related to within-levels wage inequality.

[13] Following Buchinsky (1994), the quantile regression can be written as follows: The τ-th quantile of y_i conditional on the regressor $\mathbf{Z}_i$: (schooling and other regressors)

$$Q_\tau(y_i \mid \mathbf{Z}_i) = \mathbf{Z}_i'\theta(\tau)$$

pertaining to workers with similar observed characteristics at various points along the adjusted wage scale due to unobservable attributes. In particular, the lower and upper quantiles may be respectively thought of as workers with wages lower than and wages higher than predicted by their socio-demographic characteristics, including educational attainment. In this light, the relative positioning of workers in the conditional wage distribution been taken to reflect differences in ability (see for example Arias et al (2002) or Mwabu and Shultz (2002)). For our purposes, we use the quantile regressions results to investigate the likely heterogeneity among workers in the evolution of the skill wage premium.

Table 5 provides summary results of separate regressions for public sector and private sector wage-earners using the three benchmark specifications. To examine whether public sector wage differentials have changed over time, Table 5 pools available data into the three benchmark periods, for each country for which the relevant survey years are available.

Table 6 presents estimates of the returns to discrete educational attainment (high school and at least some college education) for each country. Like the preceding table, Table 6 pools the data into the three benchmark periods, to test whether returns to schooling levels have risen over time.

Table 7 reports returns to college using additional information on family backgrounds and discusses the robustness of result with respect to endogeneity following the framework that will be presented in Section 6.

Finally, Table 8 uses the coefficient estimates of years of schooling from the regression results of each of the three specifications reported in Table 1 to run a cross-country regression of country-specific returns to schooling on a number of country-level explanatory variables, including the private sector share of GDP, FDI flows (in percent of GDP), the EBRD index of enterprise reform, and other macroeconomic indicators.

5. Results: OLS Earnings Regressions

Estimates of country-year specific returns to one year of schooling are reported in Table 1. The basic specification shows a wide range of values: returns can be as low as 2.8% for Russia at the beginning of the transition (1991) or as high as 11.1% for Hungary in late transitions years (2002). The richer specification usually implies, as expected, lower returns with drops that are almost always statistically significant.

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where $\theta(\tau)$ is the slope of the quantile line on the τ-th quantile conditional quantile of y_i. The regression residuals are such that τ are below the regression line and (1- τ) are above the regression line. For simplicity, we restrict the analysis to τ=0.1, 0.2, 0.5 (or median regression), 0.7, and 0.9.

Looking at cross-country comparisons, we can first summarize the information by focusing on the last year available for each country (2002 or 2001). A first group (Hungary and Poland) has quite high returns: more than 10% in the basic specification and about 7% in the richer. On a second group (Bulgaria, Latvia, Slovenia, and Russia) we estimate lower returns: about 7-8% in the basic specification and 4-5% in the richer specification except Latvia that experiences a larger drop to 2.5%. Finally we observe a third group (Czech Republic and Slovak Republic) with returns ranging from about 6% to about 3%. In early transition years, the mid 1990s, this grouping is roughly respected while we observe some variations in some specific years. These rates of return are broadly comparable with those of high income economies (about 7.4%) and OECD economies (7.5%), according to the most recent global survey of returns to schooling (Psacharopoulos and Patrinos 2004). A recent collective volume on returns to schooling in Western Europe utilizing comparable methodology also suggest that returns to schooling are in the 5-11% range (Colm and others 2001).

Country-specific variations over time are in the order of a couple of percentage points, except Bulgaria, Latvia and Slovak Republic that experience quite stable returns. For example, returns for Bulgaria vary from 4.7% in 1992 to 7.2% in 2002 and for Poland from 6.0% in 1991 to 10.6% in 2002. In Hungary and Russia, we estimate greater variation over time, although they are also the two countries for which we have a longer time series. In particular, Hungary is the only country where we have enough information to examine the pre-transition period. Hungary and Russia are also the only two countries that show clearly increasing trends in estimated returns. Using the results of the basic specification in Table 1, returns to schooling have increased in Hungary from about 5.6% to 11% between 1986 and 2002 and in Russia from 2.8% to 7.4% between 1991 and 2002.These are substantial increases, considering that globally, returns to schoolings have generally been gradually declining (owing in large part to an increase in the supply of schooling) (Psacharopoulos and Patrinos 2004 and Psacharopoulos 2004). The recent volume on returns to schooling in Western Europe in the 1980s and in the 1990s also suggests no discernable trend in returns to schooling and, at best, only a marginal increase in returns to schooling (Colm and others 2001).

How do these estimates compare with existing studies of returns to each year of schooling? As noted previously, few studies cover sufficiently long time periods within each country. Those that do, however, appear to have estimates broadly similar to ours (see Appendix Table 1). In the case of Hungary, for example, we estimate that returns increase from about 5.6 in 1986 to over 11 percent in 2002. This is broadly consistent with Campos and Jolliffe (2004) who use five years of data from the Wage and Earnings Survey (WES) and conclude that returns increased from 6.1 percent in 1986 to 11.7 percent in 1998 (or 10.4 with selection correction). In Russia, the World Bank (2003) shows that returns to schooling increased from about 3.4 and 3.8 percent in 1992 for men and women, respectively, to 6.8 and 7.6 percent in 2000. This compares well with our estimated aggregated increase from 3.8 percent to 8.3 percent over the same period. Returns to schooling in Poland from the pre-transition period through 1996 estimated by Rutkowski (1997) are also similar to our estimates.[14]

To further investigate this issue, Table 2 reports estimates obtained by pooling data over the pre-, early, and late transition periods. Test for structural breaks, that is, tests for difference in the estimated returns over the different periods, are reported in the third line for each country, below the estimated coefficients and the standard errors. Equality of returns over-time is rejected under all specifications only for Hungary and Russia, confirming the indication of an increasing trend found in Table 1.

Next, we specifically estimate these trends in returns to education. The estimate of the coefficient on the interaction between years of schooling and the linear and quadratic trends are reported in Table 3. In the basic specification, we find that only Hungary, Russia and Slovenia have an increasing trend using both linear and quadratic trend specifications. However, when we implement the richer specification, only Russia is observed to experience an increasing trend both in the linear and quadratic trend specification. For Slovenia it is significant only under the quadratic trend. On the other hand, Poland has a significant trend for the linear trend model for both basic and richer specification, but not with the quadratic trend. The lack of robustness in Slovenia's regression results is consistent with the insignificant structural breaks reported in Table 2.

Quantile regression results, reported in Table 4, help explain the weak trend in returns in Hungary. The results demonstrate that average results in Hungary are driven by a combination of stable or decreasing trend in returns in the lower end of the conditional wage distribution and an increasing trend in returns in the upper end of the conditional wage distribution. For Russia, in contrast, the increases in returns to education are roughly proportional over the conditional wage distribution. The quantile regression results also provide evidence of heterogeneity in returns to education across countries as well as within countries. For example, in the Czech Republic, Latvia, and Poland, regressions results based on both the basic and richer specifications generally indicate that workers at the lower end of the conditional wage distribution have lower returns to education. On the other hand, workers at the lower end of the distribution in Russia and Bulgaria have higher returns, although this relationship appears to have reversed over time in Bulgaria. The other countries in the sample exhibit roughly constant returns over the conditional wage distribution.

Table 5 provides further evidence that the documented increases in returns to education in Hungary and Russia-- as well as the generally weaker evidence of increasing

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[14] See also Galasi and Varga (2002) for a similar estimated magnitude of increase in returns to schooling in Poland between 1992 and 2000.

returns in other countries--appear to be robust to further disaggregations, in particular, the separate regressions for public and private sector workers. Table 5 also provides supplementary evidence on what may be driving the relatively weak estimated increase in aggregated returns—in particular, the possibility of offsetting developments in returns, depending on the sector of employment. Where the increase in private sector returns is modest, the secondary influence on public sector wage may also be modest, such as in the case of Bulgaria (Jones and Simon, 2004).

The regression results show that returns to education have increased for both public and private sector workers over time in Hungary and Russia. In contrast, in other countries returns appear to have fallen or remained stable in one sector while increasing in the other. In Poland, for example, returns to education increased among workers in the private sector, while falling in the public sector; in Latvia, returns among private sector workers appear to have fallen while increasing sharply in the public sector, leading to some convergence in returns over time. Similarly, public sector workers in the Czech Republic also experienced sharper increases in returns to education. Not surprisingly, the public sector wage differential was significant in the early transition period in Latvia and the Czech Republic, but not in the late transition period. In general, the public sector wage differential is not significant in the late transition period for all the countries in the sample, except for the Slovak Republic where however, due to data limitations, only the basic specification could be utilized.

Table 6 reports returns to schooling *levels* instead of returns to *one* year of schooling. This is the other standard specification used in the literature to estimate returns to education. It is a slightly more general specification because returns are not constrained to be the same for each additional year of schooling. We divide schooling levels in three categories: "No High School Completed," "High School Completed," and "Some College Years Completed or More." For some countries a finer, and arguably more precise, classification is possible but we prefer to build more aggregate measures to ensure comparability across countries and across years. We use the framework that we have already applied in Tables 2 and 5: two specifications with data pooled over the pre-, early-, and late- transition periods together with tests for structural breaks.

The first interesting implication of this analysis is that the increasing trend in returns for Hungary seems mainly driven by higher returns on College or more while High School returns do not show any significant structural breaks. For Russia, results are more mixed and, in particular, equality of the coefficients on College or more is rejected under the basic specification but it is not under the richer specification. In the case of Bulgaria, contrary to the results presented in Table 2, both returns show the presence of significant structural breaks, thus confirming how a specification measuring the rate of return to *one* year of schooling masks substantial heterogeneity in returns to education. Results for the other countries confirm the picture of lack of significant structural breaks even if, in general, returns to College or more are increasing more than returns to High School. The ranking of returns across countries are roughly respected with the exception of Poland. Under the richer specification, Poland yields relatively low returns while in Table 1 and 2, it shows relatively high returns under any specification.

6. Structural Interpretation of Results and Endogeneity

An extensive literature has emerged to investigate the consistency of returns to schooling estimated in the context of standard earnings regressions as the ones presented so far in the paper. Card (2001) summarizes the issue using a simple model of optimal investment in homogenous schooling based on Becker (1967). The conclusion is that OLS is biased unless the following assumptions are satisfied:

Assumption 1*: Homogenous marginal return to schooling in the population*

Assumption 2*: Homogenous marginal cost of schooling*

Under the model and the previous assumptions, the coefficient γ in equation (1) is the rate at which marginal benefit and marginal cost of schooling are equal for individuals choosing years of schooling to maximize lifecycle utility. Heckman, Lochner and Todd (2005) work under these assumptions to characterize carefully the structural interpretation of the γ coefficient. In their model, based on Mincer (1958 and 1974), individuals choose the schooling level to maximize the present value of earnings. In a compensating differential framework with:

Assumption 3*: Perfect certainty about the present value of future earnings*

The coefficient γ represents the discount rate that guarantees indifference (equilibrium) between the various schooling levels, i.e. the discount rate that equates the earnings streams associated with each of them.

To evaluate the consistency of the estimates obtained so far and to support their interpretation, it is interesting to ask how plausible these 3 assumptions are in the particular application we are considering. With respect to Assumption 1, we do not see any reason why its validity should be stronger or weaker in a transition economy context as compared to industrialized countries where much of the literature has focused to date (typically, the U.S.). Given the more centralized wage-setting mechanism before the transition and in the early transition stage, it is possible that the ability bias is milder. However, this is a speculation difficult to judge with the data available. With respect to Assumption 2, we believe its plausibility is stronger in transition economies under consideration because the direct cost of education is very low and the opportunity cost is generally lower than in industrialized countries.[15] Finally, with respect to Assumption 3, we consider perfect certainty as a good approximation of the pre-transition environment while in the post-transition phase, its change is potentially very important. Still, violation of Assumption 3 is generally not considered the main problem in the literature, although Heckman, Lochner and Todd (2005) show that this may well not be the case for the U.S.

Most of the literature has focused on violations of the first two assumptions. If there is heterogeneity in marginal benefits and costs and if this heterogeneity is correlated with schooling, then years of schooling in equation (1) are endogenous and OLS is an inconsistent estimator of the average return to one year of schooling.

One of the solutions most frequently employed to solve this endogeneity problem is to perform an instrumental variable estimation (IV). Unfortunately, this procedure is particularly difficult to apply here. In particular, natural experiment-type instruments are generally not comparable across countries and over time. Another solution has been to use non-parametric techniques such as matching methods. However, the amount of information we have is too limited to create credible comparison groups to evaluate the treatment. Typically, a third solution has been to develop a more general model that explicitly considers some sources of endogeneity and estimates the primitive parameters of the process that generate the observed data. This solution is probably even harder to apply here since we observe only cross-sectional data with observations on little more than wages and basic human capital variables.

Conditioning on these data limitations, we propose an incomplete solution using an additional piece of information available for a few years on selected countries: parents' education. Parents' education was one of the first instruments proposed in the literature to obtain IV estimators. However, this choice has recently been criticized because parents' education seems hardly exogenous to the wage determination process. An alternative and potentially more appropriate way to use this information is the following: first, use it as an additional regressor in the earnings regression and, second, employ it as an exclusion-restriction in the context of a control function approach (CF).[16] Adding parents' education as an additional regressor may reduce the bias due to violations of Assumptions 1 and 2. Using the control function approach has the advantage of simultaneous modeling, even in a very simple static context. Both the process of educational attainment and the process of generating wages provide consistent estimates of returns to schooling and a direct test of endogeneity. However, the consistency of the CF estimator is based on two grounds: a parametric assumption on the distribution of the unobservables responsible for the endogeneity and some exclusion restrictions. If the exclusion restrictions are not valid, all

[15] Tuition costs are usually extremely low and private education is very limited.

[16] See the seminal works by Heckman (1978 and 1979).

the identification is coming from the parametric assumption and ultimately this may be the only reason why this approach may deliver better results than an IV.[17]

With these limitations in mind, a control function approach consists of the following. Consider simply one treatment: return to College education.[18] We limit ourselves to a single treatment to have a measure more homogenous across countries. Different countries have different educational system but all countries have a quite constant definition of the College degree. Under the assumptions previously outlined, the return to college is the coefficient γ that corresponds to the dummy for college completed C_i in a standard earnings regression:

$$\log w_i = x'_i \beta + \gamma C_i + \varepsilon_i \tag{2}$$

However, the decision to complete College is endogenous. It can be modeled as an optimal decision rule based on the following latent variable:

$$C_i^* = z'_i \alpha + u_i \tag{3}$$

The latent variable represents a normalized net benefit to complete College therefore the optimal decision rule will be to complete College when the latent variable is positive. The dummy we actually observe will then be the result of:

$$C_i = I\left[C_i^* > 0\right] = I\left[u_i > -z'_i \alpha\right] \tag{4}$$

Equations (1)-(3) are a joint static model of education decision and wage determination. If the two processes are independent, that is if ε_i and u_i are uncorrelated, OLS estimation on equation (1) will provide consistent estimates. However, if ε_i and u_i are correlated, then C_i in equation (1) is endogenous and OLS is biased. Assuming ε_i and u_i are joint normal with covariance ρ, variance of the earning regression error σ and variance of the selection equation error normalized to one, the bias amounts to the second term of the following difference of conditional means:[19]

$$E\left[\log w_i \mid x_i, z_i, C_i = 1\right] - E\left[\log w_i \mid x_i, z_i, C_i = 0\right] = \gamma + \rho\sigma\left[\frac{\phi(z'_i \alpha)}{\Phi(z'_i \alpha)(1-\Phi(z'_i \alpha))}\right] \tag{5}$$

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[17] Vella and Verbeek (1999) illustrate this point: IV essentially amounts to estimating the original wage regression adding the OLS residual of the selection equation while CF amounts to adding the generalized probit residual of the same selection equation.

[18] With the same approach is possible to allow for more then one treatment by changing the first stage of the procedure. For example an ordered probit can be used to study returns to different levels of education (see for example Bratti et al 2005 and Blundell et al 2004).

[19] The result is obtained from standard properties of the truncated bivariate normal. See for example Theorem 22.5 in Greene (2003).

where usual notation is used for the *pdf* and *cdf* of the standard normal. Equation (5) shows that under the model OLS is unbiased if and only if $\rho=0$, that is if the college decision is exogenous. A test for $\rho=0$ will therefore constitutes an endogeneity test.

The identification of the parameters (β, γ, α, ρ, σ) is warranted by knowledge of (w_i, x_i, z_i). However, if $x_i=z_i$ the identification is obtained only through the nonlinearity introduced by the normality assumption. It is therefore common practice to introduce exclusion restrictions that can make the identification less dependable from functional form. We use parents' education as an exclusion restriction. The assumption is that parents' education has a significant impact on college decision but a limited impact on the wage process. In terms of the previous notation, parents' education is included in the z_i vector but not in the x_i vector. A standard F-test provides an assessment of this assumption.

The estimation of the previous model may be obtained by a two-step procedure (Heckman 1978): first, a probit on the schooling decision and then an OLS on a regression corrected for selection. To increase efficiency, we implement a full information maximum likelihood estimation of the model.

Results are presented in Table 7. The first column reports the OLS estimates of return to college using the richer specification. This is the reference point to evaluate the robustness of the results obtained so far in the paper. The second column reports the OLS estimates of returns to college using the same specification plus parent's education. Parents' education is simply summarized by father's years of schooling completed.[20] This choice is motivated by data availability and to maintain the same specification across countries and over time. The P-value of a test for the significance of father's education in the earnings regression is provided below the point estimate and standard error of the coefficient. The test is useful to evaluate our exclusion-restrictions: lack of significance in the earnings regression and significance in the selection equation of the control function approach, reported on the same row of the third column, means not rejection of the exclusion-restrictions. The third column reports returns to college estimated by CF: on top of the previous exclusion-restriction test, an endogeneity test is provided on the third row, i.e. a Likelihood-Ratio test for the null $\rho=0$.

The results are ambiguous with respect to what is the more appropriate use of family back ground information: Father's education is only slightly more significant in the selection equation than in the earnings regression equation. Exogeneity is rejected on all the specifications except Hungary in 1988 and 1992. These results confirm that OLS

[20] We have chosen father's education instead of mother's education (or some other indicator based on both) because for a few country-year surveys, only father's education was available.

estimates are biased, though we recognize that our proposed solution only partially addresses the problem.

Comparisons of changes in returns to education over time between Hungary and Poland—the only two countries on which we have family background information on more than one year—do not suggest a strong trend in returns to college. However, returns to college in Hungary increase from 16.1% in the first pre-transition year available (1988) to more than 50% in the last post-transition year available (1993) and in Poland they increase from more than 50% in 1991 to more than 70% in 1993.

The comparison across countries is, however, sensitive to the estimation method. For example, comparing Hungary and Poland in the same years with OLS suggest higher returns in Hungary while CF implies higher returns in Poland. For Russia and Slovak Republic, instead, both OLS and CF show the same ranking and the ranking is in agreement with what we have found in the previous sections of the paper: Russia has higher returns to schooling than the Slovak Republic in late transition years.

Overall results show that OLS is likely to be downward biased. However, this inference is obtained by CF estimates which are sensitive to the exclusion restrictions we have adopted. Our conclusion from this section is therefore a word of caution in interpreting the results. Even if we do not reject the presence of endogeneity, data limitations do not allow us to assess if the lack of trend (or the weak trend) that we find in the first part of the paper is due in part to the inconsistency of our estimates.

7. Cross-Country Differences in Returns to Schooling

The results obtained so far are a description of returns to schooling over-time across transition economies using a homogenous and comparable sample of individuals. An interesting additional step is to correlate cross-country changes in returns with some characteristics of the economies under consideration. An empirical model for such an analysis is presented in Fleisher, Sabirianova and Wang (2005), although their meta-analysis uses estimated returns drawn from various studies utilizing different datasets and estimation strategies. In contrast, we have at our disposal estimated returns to education drawn from comparable data and based on a common empirical framework. Building on their approach, we use the returns obtained under the three different specifications reported in the previous section, to estimate the following modified version of their extended model:

$$r_{it} = \gamma_0 + \gamma_1 t + \gamma_2 C_i + \gamma_3 R_{it} + M'_{it}\gamma_4 + \varepsilon_{it} \qquad (6)$$

where: (i,t) denotes country i in time t; r denotes the point estimate of the return to one year of schooling; t is the time trend; C is the country fixed effect; R is a measure of reform progress; M is a vector of macroeconomic controls. To define R we again follow Fleisher, Sabirianova and Wang (2005)[21] using different measures of reform. The first is the EBRD Index of Enterprise Reform, the second is the private sector share in the GDP and the third is the Foreign Direct Investment (FDI) share in the GDP. We use these three indicators separately in there different regressions and jointly in a fourth one. The macroeconomic controls are real GDP growth and the rate of inflation computed from the CPI index.[22]

The results are reported in Table 8. Only the coefficient estimates of the measures of reform are reported, while the time trend (always slightly positive), the country fixed-effect and the macroeconomic controls are omitted. The top panel reports Ordinary Least Squares (OLS) regressions: they generate coefficients that are, in general, not significantly different from zero. However, we suspect that this is due to an efficiency problem, given the relatively low number of observations. Indeed, results from Weighted Least Squares (WLS) generate significant effects of all the three reform progress variables in the basic specification. In the richer specification, instead, only the impact of the FDI share of GDP is significantly different from zero. The WLS estimates are obtained using as weight the standard error of the estimated return to one year of schooling.

Our results indicate that the usual positive correlation between the speed of structural reforms and returns to schooling is not very robust. While two indicators out of three (the EBRD Index of Enterprise Reform and the Private Sector Share of GDP) are positively correlated with the estimated returns to one of year of schooling, their impact is not significant when country fixed-effects, a time trend and macroeconomic controls are taken into account. Moreover, the third indicator (FDI share of GDP) has a relatively robust negative impact, although this may, in part, be due to selection. As our measures of reform are a standard description of liberalization that should affect the labor market, we find these results consistent with our previous evidence of a lack of robustly increasing returns following market liberalization.

One important limitation of the analysis is the absence of a specific reform measure of labor market institutions. Existing indexes of reform typically do not include a specific index of labor market reform. The one comparable cross-country index of labor market reform (and available over a sufficiently long time period) is the index of wage liberalization in the same dataset compiled by Fleisher, Sabirianova and Wang (2005). This

[21] We are very grateful to the authors and, in particular, to Klara Sabirianova Peter, for sharing the data with us: we have used their measure of reform progress and their macroeconomic controls foe all countries except Latvia which is not present in their sample. For Latvia we have constructed the same variables using the primary sources (EBRD and WDI).

[22] It seems quite natural to introduce in the specification also an indicator for wage liberalization. However, for the period and countries under consideration, the indicator does not show enough variability to provide credible estimates. We have attempted a specification with a dummy=1 when the highest score in the wage liberalization indicator is obtained but also in this case the variability is limited and the variable does not show a significant impact.

wage liberalization index, however, had limited variability. Although the evolution of rates of returns to schooling across countries is positively and significant related to this index, consistent with expectations, the results of further analyses suggest that the initial results may be driven in large part by the outliers.

8. Concluding Observations

In this paper we performed a more complete and robust analysis of the assertion that labor market liberalization leads to higher returns to schooling. In general, we are unable to find robust empirical evidence in support of this statement as the main empirical results are sensitive to the specification and estimation methods. For two countries that we observe over a long time span (Hungary and Russia), we do find a robust, increasing trend in returns to schooling. We also find large cross-country variations in the level of returns to schooling, although these rates of return are all generally comparable with those of high-income economies. The countries in our sample can be classified into three broad groups: a "high" returns group (Hungary and Poland), a "medium" returns group (Bulgaria, Latvia, Slovenia and Russia) and a "low" returns group (Czech Republic and Slovak Republic). These rankings are generally robust to the specification and the estimation method.

The results may at first seem surprising, given the documented increases in returns to education as well as the rise in wage inequality over the transition period. We argue, however, that these findings are consistent with a few recent papers in the literature. In particular, at the beginning of this paper, we provided a survey of studies suggesting that the largest increases in returns to education took place in the early transition period, with largely stable returns after 1993/1994. Except for Hungary, the period beginning from the early transition is precisely the period covered by the analysis in this paper. In addition, data on earnings inequality collected by the United Nations Economic Commission for Europe (UNECE) indicate the relative stability of earnings inequality in the Czech Republic, Latvia, and Slovenia—as measured by the Gini coefficient—since the early transition period (UNECE 2004: p. 167).[23] Furthermore, many of the countries in our sample are European transition economies, with Russia being the lone country from the Commonwealth of Independent States (CIS). The absence of a robust increasing trend in returns to schooling—other than in Russia and Hungary—is arguably consistent with differences in levels of wage inequality among sub-groups of transition economies. In particular, wage inequality in European transition economies is relatively low and comparable to OECD averages while wage inequality in CIS countries is much higher and comparable to developing country averages (see for example World Bank 2005).

[23] For example, in the Czech Republic, the Gini coefficient for earnings was 0.26 in 1994; in 2001, it was 0.27. In contrast, the Gini coefficient in Russia over this same period increased from 0.44 to 0.52.

What could be driving the relatively weak documented increase in returns to education? We reexamine some of the hypotheses put forward at the beginning of this paper and speculate on how some of our results relate to them.

Economic reform and the speed and uniformity of transition. The main hypothesis in the literature on wages and human capital in transition economies focuses on the role of market liberalization and competition in raising the estimated returns to schooling. In practice, however, the speed of transition varies from country to country. Within each country, firms and sectors will continue to function alongside each other while managed differently and differently responsive to the new market environment.

While this paper is unable to establish a systematic link between returns to schooling and the speed of economic reform, the heterogeneity in returns to schooling *within* countries suggests that non-competitive forces may have continued to determine, in large part, wage levels and their distribution. This is consistent with other studies in the literature which have found that while returns to schooling have increased, non-market forces remain powerful determinants of wages (e.g., Jones and Simon 2004). As a result, the transition process is never quite uniform, even for workers and firms within the same country. We show, for example, how returns to schooling may have evolved differently for workers at various points along the conditional wage distribution. We also provide evidence that returns to schooling may have risen at different speeds across sectors. The sample of countries covered in this paper suggests a full spectrum of possibilities—namely, that private sector returns may have risen more quickly, more slowly, or proportionally, in relation to public sector returns. In some cases, these returns may have fallen as well. In countries where returns to schooling in the private sector have decreased, or increased only modestly, or where the private sector share of employment remains small, we also expect that the secondary impact of rising private sector returns on public sector wage distribution may be just as modest.

The role of educational systems. It has been hypothesized that decreasing returns to skills acquired during the socialist period may offset the expected positive impact of market liberalization on returns to education. Unfortunately, the relatively small sample size of the ISSP surveys does not allow us to compare the changes in the returns to education over time by education cohorts (i.e., comparing the returns to education among workers educated under the new system with older workers who obtained their education under the socialist system). In addition to the small sample size, there is also the difficulty of finding a sufficiently long time period that allows post-transition human capital to accumulate returns.

Nevertheless, education systems in transition economies may not have adequately adjusted to the new market economy environment. A recent comprehensive review of education systems in the region by Berryman (2000) concludes that education systems focus on “memorized factual and procedural knowledge,” skills that are appropriate for planned economies but not necessarily for the demands of a new market economy. There is also some evidence that falling expenditures on education during the transition period may have significant effects on the quality of education (Campos and Jolliffe 2004). A recent report provides evidence of declining performance in education in European transition economies, including Czech Republic, the Slovak Republic, and Slovenia (Alam and others 2005).

Supply responses. Closely related to the issues of educational reform and quality of education are the variations in supply responses across countries. Adamchik (1999) suggests that in Poland, the education system continues to produce young workers with vocational education in droves, despite the fact that workers with vocational education constitute the largest pool of unemployed workers. In contrast, Campos and Jolliffe (2004) find that since 1989, a rising share of students in Hungary have chosen general education over vocational and technical education. They conclude that this is in response to a liberalized market economy, based on the assumption that under the previous system general education was “under-valued and under-supplied.”

These hypotheses are of course simplifying assumptions that are treated separately only for exposition. In reality, the rising skill wage premium, or its relative stability, may be driven by a confluence of factors and the shifting relative importance of these factors over time. Peter (2003), for example, suggests that in the case of Russia increasing returns to schooling were driven by institutional factors in the early transition period and by productivity change in the late transition period.

This is, however, substantially beyond the scope of this paper. The objective of the paper was to assemble comparable data and to subject the analysis to more comprehensive robustness tests, reflecting some advancement over the existing literature on returns to schooling in transition economies. In conclusion, we believe that the sensitivity of our results is consistent with the view that the process which is taking place in transition is very heterogeneous, not only in terms of initial conditions but also with respect to its dynamics.

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Table 1: Returns to One Year of Schooling by Country by Year

	1986	1987	1988	1989	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002
Bulgaria																	
Basic							0.047	0.052				0.053		0.050	0.049		0.072
							(0.009)	(0.009)				(0.010)		(0.010)	(0.009)		(0.011)
Basic							0.048	0.056				0.055		0.050	0.050		0.074
Balanced							(0.009)	(0.009)				(0.010)		(0.011)	(0.009)		(0.011)
Richer							0.034	0.036				0.035		0.036	0.028		0.057
							(0.012)	(0.010)				(0.015)		(0.017)	(0.010)		(0.016)
Czech R																	
Basic									0.036	0.044	0.070	0.076		0.054		0.087	0.066
									(0.008)	(0.006)	(0.007)	(0.008)		(0.006)		(0.009)	(0.010)
Basic									0.042	0.042	0.070	0.082		0.051		0.086	0.062
Balanced									(0.007)	(0.006)	(0.008)	(0.008)		(0.006)		(0.009)	(0.010)
Richer									0.028	0.033	0.048	0.066		0.030		0.063	0.034
									(0.008)	(0.008)	(0.009)	(0.009)		(0.007)		(0.011)	(0.011)
Hungary																	
Basic	0.056	0.054	0.060	0.072	0.057	0.075	0.071	0.076	0.096	0.088	0.082	0.104	0.080	0.109		0.090	0.111
	(0.004)	(0.003)	(0.004)	(0.006)	(0.006)	(0.007)	(0.005)	(0.005)	(0.007)	(0.007)	(0.005)	(0.007)	(0.008)	(0.009)		(0.012)	(0.009)
Basic			0.060	0.072				0.075	0.095	0.088	0.083		0.082	0.104		0.090	0.107
Balanced			(0.004)	(0.006)				(0.005)	(0.007)	(0.007)	(0.005)		(0.008)	(0.013)		(0.012)	(0.010)
Richer			0.044	0.053				0.051	0.056	0.062	0.050		0.053	0.106		0.063	0.070
			(0.005)	(0.008)				(0.006)	(0.007)	(0.009)	(0.006)		(0.010)	(0.017)		(0.015)	(0.010)
Latvia																	
Basic										0.067	0.047		0.053	0.086	0.065	0.080	0.078
										(0.010)	(0.008)		(0.010)	(0.010)	(0.009)	(0.009)	(0.011)
Basic										0.068	0.046		0.052	0.080	0.063	0.083	0.075
Balanced										(0.010)	(0.008)		(0.010)	(0.010)	(0.010)	(0.009)	(0.012)
Richer										0.043	0.028		0.030	0.050	0.039	0.057	0.025
										(0.012)	(0.009)		(0.013)	(0.014)	(0.015)	(0.010)	(0.013)
Poland																	
Basic						0.060	0.071	0.081	0.080	0.079	0.070	0.065	0.081	0.079		0.092	0.106
						(0.009)	(0.008)	(0.008)	(0.008)	(0.008)	(0.009)	(0.009)	(0.010)	(0.010)		(0.010)	(0.011)
Basic						0.059	0.071			0.079	0.070	0.064	0.081	0.079		0.090	
Balanced						(0.009)	(0.008)			(0.008)	(0.009)	(0.009)	(0.010)	(0.010)		(0.010)	
Richer						0.055	0.059			0.041	0.063	0.053	0.068	0.028		0.071	
						(0.013)	(0.013)			(0.010)	(0.011)	(0.014)	(0.012)	(0.012)		(0.015)	

Table 1: cont.

Russia												
Basic	0.028	0.038	0.043	0.054	0.083	0.065	0.072		0.084	0.083	0.084	0.074
	(0.003)	(0.006)	(0.008)	(0.008)	(0.011)	(0.010)	(0.011)		(0.012)	(0.011)	(0.011)	(0.013)
Basic	0.036		0.047	0.054	0.084	0.065	0.066		0.091	0.081	0.081	
Balanced	(0.006)		(0.009)	(0.009)	(0.011)	(0.011)	(0.011)		(0.013)	(0.011)	(0.011)	
Richer	0.030		0.014	0.041	0.060	0.030	0.046		0.068	0.056	0.041	
	(0.007)		(0.011)	(0.010)	(0.013)	(0.014)	(0.012)		(0.015)	(0.013)	(0.012)	
Slovak R												
Basic					0.061			0.066	0.059			0.061
					(0.007)			(0.007)	(0.006)			(0.007)
Basic									0.058			0.060
Balanced									(0.006)			(0.007)
Richer									0.028			0.029
									(0.008)			(0.007)
Slovenia												
Basic	0.063	0.058	0.088		0.095	0.117	0.099	0.089		0.081		0.082
	(0.005)	(0.007)	(0.008)		(0.008)	(0.007)	(0.009)	(0.007)		(0.007)		(0.009)
Basic	0.065	0.067	0.092		0.098	0.113	0.093	0.089		0.082		0.075
Balanced	(0.006)	(0.007)	(0.008)		(0.008)	(0.008)	(0.009)	(0.007)		(0.007)		(0.009)
Richer	0.048	0.045	0.049		0.077	0.077	0.056	0.042		0.057		0.040
	(0.007)	(0.009)	(0.011)		(0.012)	(0.011)	(0.012)	(0.010)		(0.009)		(0.009)

Note: Coefficients of the years of schooling variable in earning regressions. Dependent variables are monthly earnings. Robust standard errors in parentheses. Specifications include the following regressors. Basic: potential experience (linear and squared), dummies for male. Richer: previous regressors plus dummies for living in urban areas and married, controls for current job (dummies for occupation, public employee, working full-time, member of a trade union), controls for current family (number of members, dummy for spouse working full-time).

Table 2: Structural Breaks in Returns to One Year of Schooling by Country

	Basic Specification			Basic on Balanced Sample			Richer Specification		
Country	Pre (1986-1990)	Early (1991-1996)	Late (1997-2002)	Pre (1986-1990)	Early (1991-1996)	Late (1997-2002)	Pre (1986-1990)	Early (1991-1996)	Late (1997-2002)
Bulgaria									
Coefficient		0.050	0.056		0.052	0.057		0.034	0.044
		(0.003)	(0.005)		(0.006)	(0.005)		(0.008)	(0.008)
Test									
$\beta_{k,t} = \beta_{k,t+j}$		0.2218			0.4169			0.1967	
Czech R.									
Coefficient		0.049	0.066		0.050	0.065		0.036	0.043
		(0.010)	(0.008)		(0.004)	(0.004)		(0.005)	(0.005)
Test									
$\beta_{k,t} = \beta_{k,t+j}$		0.1319			0.1963			0.5458	
Hungary									
Coefficient	0.058	0.081	0.100	0.063	0.084	0.096	0.049	0.055	0.069
	(0.002)	(0.004)	(0.005)	(0.005)	(0.004)	(0.005)	(0.004)	(0.002)	(0.009)
Test									
$\beta_{k,t} = \beta_{k,t+j}$	0.0000	0.0028		0.0002	0.0476		0.1267	0.0870	
Latvia									
Coefficient		0.056	0.073		0.056	0.070		0.036	0.039
		(0.010)	(0.007)		(0.007)	(0.005)		(0.007)	(0.006)
Test									
$\beta_{k,t} = \beta_{k,t+j}$		0.0907			0.1610			0.7084	
Poland									
Coefficient		0.074	0.084		0.070	0.078		0.054	0.054
		(0.003)	(0.007)		(0.004)	(0.005)		(0.006)	(0.007)
Test									
$\beta_{k,t} = \beta_{k,t+j}$		0.1896			0.2641			0.9680	
Russia									
Coefficient		0.0440	0.0785		0.0547	0.0789		0.0351	0.0506
		(0.0086)	(0.0029)		(0.0040)	(0.0058)		(0.0047)	(0.0064)
Test									
$\beta_{k,t} = \beta_{k,t+j}$		0.0001			0.0126			0.0597	
Slovak R.									
Coefficient		0.061	0.062						
		(0.000)	(0.002)						
Test									
$\beta_{k,t} = \beta_{k,t+j}$		0.7744							
Slovenia									
Coefficient		0.074	0.088		0.080	0.085		0.055	0.051
		(0.009)	(0.004)		(0.003)	(0.004)		(0.004)	(0.005)
Test									
$\beta_{k,t} = \beta_{k,t+j}$		0.1510			0.6253			0.5068	

Note: The variance/covariance matrix used to compute standard errors (in parentheses) and P-values of tests is corrected for year clustering.

Table 3. Trends in Returns to One Year of Schooling

	Basic Specification		Basic on Balanced Sample		Richer Specification	
	(1)	(2)	(1)	(2)	(1)	(2)
Bulgaria						
Schooling	0.046	0.050	0.048	0.053	0.034	0.038
	(0.004)	(0.006)	(0.006)	(0.008)	(0.007)	(0.008)
Trend (linear)	0.0015	-0.0026	0.0014	-0.0032	0.0015	-0.0030
	(0.001)	(0.003)	(0.001)	(0.004)	(0.001)	(0.004)
Trend (quadratic)		0.0004		0.0005		0.0005
		(0.000)		(0.000)		(0.000)
Czech Republic						
Schooling	0.045	0.042	0.048	0.045	0.034	0.027
	(0.006)	(0.006)	(0.006)	(0.005)	(0.007)	(0.004)
Trend (linear)	0.0036	0.0058	0.0030	0.0058	0.0016	0.0084
	(0.001)	(0.005)	(0.001)	(0.005)	(0.002)	(0.005)
Trend (quadratic)		-0.0003		-0.0004		-0.0009
		(0.001)		(0.001)		(0.001)
Hungary						
Schooling	0.053	0.052	0.057	0.052	0.049	0.056
	(0.002)	(0.002)	(0.004)	(0.005)	(0.004)	(0.007)
Trend (linear)	0.0035	0.0039	0.0031	0.0045	0.0010	-0.0010
	(0.000)	(0.001)	(0.000)	(0.002)	(0.001)	(0.001)
Trend (quadratic)		0.0000		-0.0001		0.0001
		(0.000)		(0.000)		(0.000)
Latvia						
Schooling	0.058	0.061	0.056	0.060	0.034	0.038
	(0.009)	(0.009)	(0.009)	(0.008)	(0.008)	(0.008)
Trend (linear)	0.0029	-0.0002	0.0030	-0.0024	0.0015	-0.0030
	(0.002)	(0.008)	(0.002)	(0.006)	(0.002)	(0.006)
Trend (quadratic)		0.0005		0.0008		0.0007
		(0.001)		(0.001)		(0.001)
Poland						
Schooling	0.064	0.069	0.061	0.063	0.040	0.042
	(0.003)	(0.004)	(0.003)	(0.002)	(0.005)	(0.006)
Trend (linear)	0.0031	0.0000	0.0026	0.0014	0.0030	0.0020
	(0.001)	(0.002)	(0.000)	(0.001)	(0.001)	(0.002)
Trend (quadratic)		0.0003		0.0001		0.0001
		(0.000)		(0.000)		(0.000)
Russia						
Schooling	0.033	0.027	0.042	0.033	0.023	0.015
	(0.005)	(0.002)	(0.006)	(0.007)	(0.005)	(0.007)
Trend (linear)	0.0053	0.0121	0.0045	0.0097	0.0039	0.0087
	(0.001)	(0.002)	(0.001)	(0.005)	(0.001)	(0.004)
Trend (quadratic)		-0.0007		-0.0005		-0.0005
		(0.000)		(0.001)		(0.000)

Table 3: cont.

Slovak Republic						
Schooling	0.061	0.062	0.053	…	0.041	…
	(0.001)	(0.001)	(0.001)	…	(0.000)	…
Trend (linear)	0.0001	-0.0005	0.0010	…	-0.0023	…
	(0.000)	(0.001)	(0.000)	…	(0.001)	…
Trend (quadratic)		0.0001		…		…
		(0.000)		…		…
Slovenia						
Schooling	0.071	0.065	0.074	0.064	0.048	0.039
	(0.006)	(0.004)	(0.007)	(0.003)	(0.004)	(0.003)
Trend (linear)	0.0022	0.0087	0.0019	0.0105	0.0013	0.0093
	(0.001)	(0.002)	(0.001)	(0.002)	(0.001)	(0.002)
Trend (quadratic)		-0.0007		-0.0009		-0.0008
		(0.000)		(0.000)		(0.000)

Notes: Coefficient estimates by specification; robust standard errors in parentheses. The results are based on all available data for each country using the indicated specification, including year dummies and the interaction of the linear trend and the squared trend with years of schooling as indicated. The variance/covariance matrix used to compute standard errors (in parentheses) and P-values of tests is corrected for year clustering.

Table 4: Returns to One Year of Schooling: Quantile Regression

	Basic Specification						Richer Specification					
		Quantile Regressions						Quantile Regressions				
	OLS	0.1	0.2	0.5	0.7	0.9	OLS	0.1	0.2	0.5	0.7	0.9
Bulgaria												
Early (1991-1996)	0.050	0.069	0.060	0.050	0.041	0.033	0.034	0.041	0.043	0.030	0.029	0.023
	(0.003)	(0.011)	(0.008)	(0.007)	(0.008)	(0.012)	(0.001)	(0.011)	(0.011)	(0.010)	(0.010)	(0.017)
Late (1997-2002)	0.056	0.058	0.057	0.062	0.056	0.054	0.044	0.030	0.039	0.048	0.061	0.059
	(0.005)	(0.011)	(0.008)	(0.005)	(0.006)	(0.008)	(0.008)	(0.013)	(0.009)	(0.007)	(0.008)	(0.010)
Czech Republic												
Early (1991-1996)	0.049	0.042	0.039	0.050	0.060	0.060	0.036	0.022	0.023	0.037	0.050	0.052
	(0.010)	(0.006)	(0.005)	(0.003)	(0.005)	(0.007)	(0.006)	(0.006)	(0.006)	(0.005)	(0.007)	(0.010)
Late (1997-2002)	0.063	0.055	0.060	0.059	0.067	0.071	0.042	0.026	0.031	0.042	0.049	0.042
	(0.007)	(0.007)	(0.004)	(0.004)	(0.004)	(0.006)	(0.008)	(0.005)	(0.004)	(0.004)	(0.006)	(0.008)
Hungary												
Pre (1986-1990)	0.058	0.060	0.059	0.062	0.057	0.049	0.049	0.054	0.049	0.055	0.052	0.041
	(0.002)	(0.003)	(0.003)	(0.002)	(0.004)	(0.004)	(0.004)	(0.006)	(0.005)	(0.004)	(0.007)	(0.008)
Early (1991-1996)	0.081	0.075	0.079	0.073	0.076	0.081	0.055	0.053	0.052	0.056	0.059	0.052
	(0.004)	(0.004)	(0.003)	(0.002)	(0.003)	(0.004)	(0.002)	(0.004)	(0.003)	(0.003)	(0.004)	(0.005)
Late (1997-2002)	0.100	0.099	0.103	0.091	0.089	0.094	0.069	0.058	0.057	0.069	0.070	0.073
	(0.005)	(0.006)	(0.006)	(0.003)	(0.004)	(0.007)	(0.009)	(0.008)	(0.007)	(0.006)	(0.008)	(0.011)
Latvia												
Early (1991-1996)	0.056	0.051	0.055	0.052	0.056	0.062	0.036	0.032	0.040	0.030	0.037	0.040
	(0.010)	(0.013)	(0.008)	(0.009)	(0.007)	(0.012)	(0.007)	(0.014)	(0.011)	(0.011)	(0.007)	(0.015)
Late (1997-2002)	0.073	0.063	0.066	0.075	0.074	0.074	0.039	0.031	0.033	0.044	0.047	0.044
	(0.007)	(0.008)	(0.005)	(0.004)	(0.005)	(0.008)	(0.006)	(0.007)	(0.007)	(0.006)	(0.007)	(0.011)
Poland												
Early (1991-1996)	0.074	0.067	0.065	0.069	0.073	0.090	0.054	0.050	0.049	0.048	0.045	0.067
	(0.003)	(0.005)	(0.004)	(0.004)	(0.004)	(0.006)	(0.004)	(0.007)	(0.006)	(0.006)	(0.008)	(0.014)
Late (1997-2002)	0.084	0.067	0.068	0.078	0.089	0.099	0.054	0.033	0.042	0.052	0.054	0.063
	(0.007)	(0.008)	(0.003)	(0.004)	(0.005)	(0.007)	(0.010)	(0.009)	(0.007)	(0.007)	(0.010)	(0.013)

Table 4: cont.

Russia												
Early (1991-1996)	0.044	0.040	0.046	0.039	0.038	0.037	0.035	0.041	0.043	0.038	0.029	0.036
	(0.009)	(0.004)	(0.004)	(0.003)	(0.004)	(0.004)	(0.007)	(0.008)	(0.008)	(0.005)	(0.006)	(0.007)
Late (1997-2002)	0.079	0.105	0.084	0.077	0.067	0.065	0.051	0.061	0.051	0.041	0.051	0.051
	(0.003)	(0.010)	(0.007)	(0.006)	(0.007)	(0.008)	(0.006)	(0.011)	(0.009)	(0.007)	(0.010)	(0.010)
Slovak Republic												
Early (1991-1996)	0.061	0.061	0.047	0.047	0.064	0.076						
	(0.000)	(0.014)	(0.009)	(0.005)	(0.007)	(0.010)						
Late (1997-2002)	0.062	0.059	0.055	0.055	0.065	0.066	0.027	0.025	0.027	0.027	0.032	0.027
	(0.002)	(0.007)	(0.005)	(0.004)	(0.005)	(0.007)	(0.001)	(0.008)	(0.009)	(0.005)	(0.007)	(0.010)
Slovenia												
Early (1991-1996)	0.074	0.070	0.075	0.079	0.079	0.080	0.055	0.056	0.056	0.057	0.063	0.070
	(0.009)	(0.004)	(0.002)	(0.003)	(0.004)	(0.005)	(0.006)	(0.006)	(0.004)	(0.004)	(0.006)	(0.010)
Late (1997-2002)	0.088	0.094	0.091	0.090	0.090	0.085	0.051	0.056	0.054	0.050	0.049	0.051
	(0.004)	(0.007)	(0.005)	(0.003)	(0.005)	(0.008)	(0.004)	(0.007)	(0.007)	(0.004)	(0.005)	(0.008)

Notes: Coefficient of years of education. The variance/covariance matrix used to compute standard errors (in parentheses) is corrected for year clustering.

Table 5: Returns to One Year of Schooling: Private and Public Sector

	Basic		Basic-Balanced		Richer	
	Private	Public	Private	Public	Private	Public
Bulgaria						
Early (1991-1996)	0.058	0.052	0.054	0.054	0.004	0.036
	(0.029)	(0.005)	(0.036)	(0.009)	(0.069)	(0.008)
Test: $\gamma_{private} = \gamma_{public}$	0.875		1.000		0.653	
Late (1997-2002)	0.061	0.055	0.066	0.057	0.056	0.041
	(0.007)	(0.006)	(0.009)	(0.005)	(0.016)	(0.008)
Test: $\gamma_{private} = \gamma_{public}$	0.590		0.380		0.367	
Czech Republic						
Early (1991-1996)	0.060	0.047	0.060	0.046	0.043	0.031
	(0.007)	(0.009)	(0.007)	(0.010)	(0.007)	(0.004)
Test: $\gamma_{private} = \gamma_{public}$	0.094		0.105		0.012	
Late (1997-2002)	0.066	0.064	0.066	0.064	0.043	0.039
	(0.006)	(0.008)	(0.007)	(0.009)	(0.007)	(0.009)
Test: $\gamma_{private} = \gamma_{public}$	0.502		0.574		0.360	
Hungary						
Pre (1986-1990)	0.068	0.058	…	0.064	…	0.050
	(0.003)	(0.002)	…	(0.005)	…	(0.006)
Test: $\gamma_{private} = \gamma_{public}$	0.004					
Early (1991-1996)	0.081	0.082	0.079	0.086	0.047	0.056
	(0.007)	(0.005)	(0.011)	(0.004)	(0.011)	(0.003)
Test: $\gamma_{private} = \gamma_{public}$	0.896		0.453		0.438	
Late (1997-2002)	0.109	0.098	0.107	0.096	0.064	0.070
	(0.010)	(0.005)	(0.009)	(0.005)	(0.017)	(0.007)
Test: $\gamma_{private} = \gamma_{public}$	0.242		0.240		0.702	
Latvia						
Early (1991-1996)	0.083	0.047	0.088	0.047	0.049	0.033
	(0.010)	(0.017)	(0.017)	(0.017)	(0.001)	(0.012)
Test: $\gamma_{private} = \gamma_{public}$	0.000		0.000		0.153	
Late (1997-2002)	0.078	0.072	0.077	0.071	0.041	0.040
	(0.004)	(0.007)	(0.004)	(0.007)	(0.009)	(0.007)
Test: $\gamma_{private} = \gamma_{public}$	0.418		0.366		0.972	
Poland						
Early (1991-1996)	0.088	0.073	0.086	0.069	0.062	0.050
	(0.008)	(0.003)	(0.012)	(0.003)	(0.009)	(0.007)
Test: $\gamma_{private} = \gamma_{public}$	0.055		0.100		0.319	
Late (1997-2002)	0.111	0.070	0.108	0.065	0.065	0.047
	(0.006)	(0.007)	(0.006)	(0.007)	(0.014)	(0.013)
Test: $\gamma_{private} = \gamma_{public}$	0.000		0.000		0.086	

Table 5: cont.

Russia						
Early (1991-1996)	0.041	0.044	0.043	0.055	0.008	0.037
	(0.016)	(0.009)	(0.021)	(0.011)	(0.024)	(0.008)
Test: $\gamma_{private} = \gamma_{public}$	0.868		0.658		0.270	
Late (1997-2002)	0.083	0.080	0.080	0.083	0.052	0.051
	(0.004)	(0.004)	(0.006)	(0.006)	(0.007)	(0.006)
Test: $\gamma_{private} = \gamma_{public}$	0.739		0.731		0.963	
Slovak Republic						
Early (1991-1996)	0.084	0.060	…	…	…	…
	(0.000)	(0.000)	…	…	…	…
Test: $\gamma_{private} = \gamma_{public}$	0.264					
Late (1997-2002)	0.068	0.061	…	…	…	…
	(0.004)	(0.001)	…	…	…	…
Test: $\gamma_{private} = \gamma_{public}$	0.019					
Slovenia						
Early (1991-1996)	0.066	0.075	0.072	0.081	0.041	0.056
	(0.028)	(0.007)	(0.033)	(0.008)	(0.023)	(0.005)
Test: $\gamma_{private} = \gamma_{public}$	0.678		0.727		0.474	
Late (1997-2002)	0.084	0.088	0.082	0.086	0.044	0.054
	(0.006)	(0.004)	(0.007)	(0.005)	(0.007)	(0.004)
Test: $\gamma_{private} = \gamma_{public}$	0.356		0.566		0.147	

Note: The variance/covariance matrix used to compute standard errors (in parentheses) and P-values of tests is corrected for year clustering.

Table 6: Returns to Schooling Levels By Country

	Basic Specification			Basic on Balanced Sample			Richer Specification		
Country	Pre (1986-1990)	Early (1991-1996)	Late (1997-2002)	Pre (1986-1990)	Early (1991-1996)	Late (1997-2002)	Pre (1986-1990)	Early (1991-1996)	Late (1997-2002)
Bulgaria									
High School		0.146	0.314		0.141	0.311		0.044	0.221
		(0.007)	(0.041)		(0.049)	(0.048)		(0.045)	(0.049)
Some College+		0.332	0.541		0.328	0.546		0.122	0.399
		(0.018)	(0.018)		(0.052)	(0.050)		(0.060)	(0.060)
Test: $\beta_{k,t} = \beta_{k,t+j}$									
High School		0.0000			0.0001			0.0028	
Some College+		0.0000			0.0000			0.0000	
Czech Rep									
High School		0.102	0.187		0.100	0.195		0.068	0.120
		(0.027)	(0.021)		(0.024)	(0.023)		(0.024)	(0.022)
Some College+		0.311	0.435		0.321	0.432		0.187	0.247
		(0.064)	(0.027)		(0.029)	(0.029)		(0.035)	(0.032)
Test: $\beta_{k,t} = \beta_{k,t+j}$									
High School		0.0050			0.0031			0.1270	
Some College+		0.0438			0.0796			0.2478	
Hungary									
High School	0.210	0.252	0.290	0.260	0.290	0.284	0.176	0.151	0.169
	(0.020)	(0.028)	(0.020)	(0.058)	(0.019)	(0.028)	(0.064)	(0.022)	(0.027)
Some College+	0.411	0.585	0.750	0.464	0.626	0.712	0.312	0.341	0.487
	(0.018)	(0.044)	(0.040)	(0.041)	(0.037)	(0.040)	(0.049)	(0.020)	(0.052)
Test: $\beta_{k,t} = \beta_{k,t+j}$									
High School	0.1933	0.2468		0.5352	0.8577		0.6325	0.5776	
Some College+	0.0001	0.0034		0.0005	0.0951		0.4805	0.0050	
Latvia									
High School		0.164	0.200		0.151	0.185		0.094	0.129
		(0.016)	(0.027)		(0.064)	(0.037)		(0.059)	(0.034)
Some College+		0.400	0.461		0.376	0.444		0.210	0.230
		(0.053)	(0.046)		(0.059)	(0.037)		(0.058)	(0.039)
Test: $\beta_{k,t} = \beta_{k,t+j}$									
High School		0.1984			0.1745			0.1050	
Some College+		0.3077			0.2263			0.5858	

Table 6: cont.

Poland						
High School	0.258	0.239	0.238	0.227	0.143	0.115
	(0.016)	(0.026)	(0.022)	(0.027)	(0.024)	(0.028)
Some College+	0.458	0.540	0.427	0.492	0.270	0.266
	(0.027)	(0.056)	(0.028)	(0.033)	(0.036)	(0.042)
Test: $\beta_{k,t} = \beta_{k,t+j}$						
High School	0.5253		0.7401		0.5587	
Some College+	0.1607		0.2140		0.9569	
Russia						
High School	0.143	0.296	0.182	0.319	0.104	0.158
	(0.055)	(0.040)	(0.038)	(0.050)	(0.037)	(0.046)
Some College+	0.367	0.597	0.441	0.611	0.273	0.348
	(0.087)	(0.028)	(0.041)	(0.051)	(0.044)	(0.051)
Test: $\beta_{k,t} = \beta_{k,t+j}$						
High School	0.0179		0.0062		0.2468	
Some College+	0.0081		0.0248		0.3459	
Slovak Rep						
High School	0.093	0.112				
	(0.000)	(0.035)				
Some College+	0.333	0.345				
	(0.000)	(0.012)				
Test: $\beta_{k,t} = \beta_{k,t+j}$						
High School	0.5676					
Some College+	0.3156					
Slovenia						
High School	0.254	0.316	0.267	0.298	0.183	0.204
	(0.031)	(0.049)	(0.022)	(0.029)	(0.023)	(0.029)
Some College+	0.649	0.669	0.654	0.637	0.424	0.342
	(0.044)	(0.031)	(0.028)	(0.033)	(0.036)	(0.040)
Test: $\beta_{k,t} = \beta_{k,t+j}$						
High School	0.2418		0.5293		0.6004	
Some College+	0.6903		0.7724		0.0999	

Note: The test reports the null hypothesis and the P-values. The variance/covariance matrix used to compute standard errors (in parentheses) and P-values is corrected for year clustering.

Table 7: Robustness to Schooling Endogeneity: Returns to Some College or More by Available Countries

	OLS	OLS with Father's Schooling	Control Function
Hungary: 1988			
Return	0.175	0.174	0.161
	(0.035)	(0.035)	(0.122)
Tests (P-value):			
Father's Schooling		0.882	0.000
Endogeneity			0.903
Hungary: 1989			
Return	0.139	0.125	0.339
	(0.043)	(0.043)	(0.122)
Tests (P-value):			
Father's Schooling		0.067	0.000
Endogeneity			0.097
Hungary: 1992			
Return	0.164	0.140	0.274
	(0.064)	(0.064)	(0.092)
Tests (P-value):			
Father's Schooling		0.029	0.000
Endogeneity			0.127
Hungary: 1993			
Return	0.204	0.191	0.584
	(0.042)	(0.042)	(0.085)
Tests (P-value):			
Father's Schooling		0.046	0.000
Endogeneity			0.001
Poland: 1991			
Return	0.226	0.197	0.552
	(0.068)	(0.068)	(0.120)
Tests (P-value):			
Father's Schooling		0.004	0.000
Endogeneity			0.006
Poland: 1992			
Return	0.133	0.117	0.320
	(0.061)	(0.061)	(0.107)
Tests (P-value):			
Father's Schooling		0.011	0.000
Endogeneity			0.048
Poland: 1993			
Return	0.157	0.142	0.731
	(0.062)	(0.061)	(0.100)
Tests (P-value):			
Father's Schooling		0.000	0.000
Endogeneity			0.000

Table 7: cont.

Russia: 1999			
Return	0.321	0.258	1.082
	(0.070)	(0.070)	(0.167)
Tests (P-value):			
Father's Schooling		0.001	0.000
Endogeneity			0.000
Slovak R: 1999			
Return	0.060	0.064	0.521
	(0.037)	(0.037)	(0.082)
Tests (P-value):			
Father's Schooling		0.000	0.000
Endogeneity			0.000

Note: Coefficients of the years of schooling variable in the "richer specification" earning regressions. The test reports the p-values for significance of father's schooling in the earning regression (second column) or in the selection equation (third column) and for endogeneity of the schooling variable (third column).

Table 8: Fixed Effects Cross-country Regression on Returns to One Year of Schooling

Returns from:	Basic Specification				Basic on Balanced Sample				Richer Specification			
	N = 67				N = 56				N = 56			
OLS:												
Enterprise	0.007			0.003	0.001			-0.001	0.001			-0.003
Reform	(0.006)			(0.006)	(0.006)			(0.007)	(0.007)			(0.008)
Private		0.051		0.041		0.034		0.028		0.035		0.032
Sector share		(0.022)		(0.024)		(0.026)		(0.030)		(0.031)		(0.036)
FDI			-0.137	-0.124			-0.154	-0.140			-0.142	-0.126
share			(0.063)	(0.062)			(0.067)	(0.069)			(0.081)	(0.084)
WLS:												
Enterprise	0.009			0.003	0.004			0.000	0.000			-0.003
Reform	(0.004)			(0.004)	(0.004)			(0.005)	(0.005)			(0.006)
Private		0.063		0.052		0.041		0.029		0.024		0.017
Sector share		(0.012)		(0.013)		(0.016)		(0.019)		(0.019)		(0.022)
FDI			-0.149	-0.111			-0.134	-0.114			-0.116	-0.106
share			(0.040)	(0.041)			(0.043)	(0.044)			(0.050)	(0.052)

Note: Dependent variable: country-year specific returns to one year of schooling (see Table 1). Regressors: Enterprise Reform is the EBRD index of enterprise reform, Private Sector share is the private sector share of GDP; FDI share is the net inflows of Foreign Direct Investment as share of GDP. Each specification includes a time trend, country fixed effects and macroeconomic controls (see Equation 6). Standard errors in parentheses. WLS means Weighted Least Squares estimates where the weight is the standard error of the estimated return to one year of schooling.

Appendix Table 1. The Returns to Schooling: Selected Point Estimates
(Beta coefficients from OLS Mincer regressions)

Country	Source	Year	All	Men	Women	Controls
Bulgaria	Rutkoswki (2001)	1997	0.065	0.062	0.068	Experience, experience-squared
Czech Republic	Munich, Svenar, Terrell (1999)	1955-90		0.017	0.036	Experience, experience-squared, industry, location, income definition, trend interaction
		1991-96		0.022	0.046	Experience, experience-squared, industry, location, income definition, trend interaction
	Munich, Svenar, Terrell (1999)	1989		0.027	0.038	Experience, experience-squared, industry, location, income definition
		1996		0.058	0.070	Experience, experience-squared, industry, location, income definition
	Chase (1998)	1984		0.024	0.042	Experience, experience-squared
		1993		0.052	0.058	Experience, experience-squared
		1984		0.030	0.048	Experience, experience-squared, industry, region
		1993		0.049	0.052	Experience, experience-squared, industry, region
	Flanagan (1998)	1988	0.043	0.037	0.051	Experience, experience-squared
		1996	0.057	0.045	0.070	Experience, experience-squared
	Vecernik (1995)	1988		0.040	0.057	Experience, experience-squared
		1992		0.053	0.067	Experience, experience-squared
	Vecernik (2001)	1988	0.044	0.040	0.051	Experience, experience-squared
		1992	0.069	0.061	0.080	Experience, experience-squared
		1996	0.088	0.083	0.094	Experience, experience-squared
	Filer, Jurajda, Planovsky (1999)	1995		0.081		Age, age-squared, industry, region, firm size, and the employer's ownership type
		1996		0.085		Age, age-squared, industry, region, firm size, and the employer's ownership type
		1997		0.090		Age, age-squared, industry, region, firm size, and the employer's ownership type
Hungary	Campos and Jolliffe (2002)	1986	0.062			Experience, experience-squared, gender, firm size, industry, region
		1989	0.074			Experience, experience-squared, gender, firm size, industry, region
		1992	0.095			Experience, experience-squared, gender, firm size, industry, region
		1995	0.109			Experience, experience-squared, gender, firm size, industry, region
		1998	0.112			Experience, experience-squared, gender, firm size, industry, region
	Campos and Jolliffe (2004)	1986	0.061			Experience, experience-squared, gender, firm size, industry, region
		1989	0.078			Experience, experience-squared, gender, firm size, industry, region
		1992	0.096			Experience, experience-squared, gender, firm size, industry, region
		1995	0.112			Experience, experience-squared, gender, firm size, industry, region
		1998	0.117			Experience, experience-squared, gender, firm size, industry, region
	Lorenz and Wagner (1993)	1987		0.039		Experience, experience-squared, outlier dummy
	Varga and Galasi (2002)	1992	0.094	0.091	0.096	Experience, experience-squared
		2000	0.104	0.105	0.105	Experience, experience-squared
	Varga and Galasi (2002)	1992	0.040	0.040	0.039	Experience, experience-squared, occupation, industry, region, sector

Appendix Table 1. The Returns to Schooling: Selected Point Estimates
(Beta coefficients from OLS Mincer regressions)

Country	Source	Year	All	Men	Women	Controls
Poland	Rutkoswki (1996)	1987	0.050			Experience, experience-squared, gender
		1992	0.079			Experience, experience-squared, gender
		1993	0.073			Experience, experience-squared, gender
	Rutkowski (1997)	1987	0.050			Experience, experience-squared, gender
		1992	0.079			Experience, experience-squared, gender
		1995	0.070			Experience, experience-squared, gender
		1996	0.078			Experience, experience-squared, gender
	Rutkoswki (2001)	1996	0.073	0.078	0.067	Experience, experience-squared, gender
	Lorenz and Wagner (1990)	1986		0.024		Experience, experience-squared, outlier dummy
Russia	Brainerd (1998)	1991		0.031	0.054	Experience, experience-squared, region, month
		1993		0.066	0.074	Experience, experience-squared, region, month
		1994		0.067	0.096	Experience, experience-squared, region, month
	Nesterova and Sabirianova (1998)	1994	0.078			Experience, experience-squared, gender, tenure, tenure-squared
		1995	0.064			Experience, experience-squared, gender, tenure, tenure-squared
		1996	0.061			Experience, experience-squared, gender, tenure, tenure-squared
	Vernon (2002)	1992		0.052	0.054	Experience, experience-squared, region, marital status, employment, hours, arrears, etc.
		1994		0.073	0.091	Experience, experience-squared, region, marital status, employment, hours, arrears, etc.
		1996		0.062	0.077	Experience, experience-squared, region, marital status, employment, hours, arrears, etc.
		1998		0.084	0.102	Experience, experience-squared, region, marital status, employment, hours, arrears, etc.
		2000		0.085	0.103	Experience, experience-squared, region, marital status, employment, hours, arrears, etc.
	World Bank (2003)	1992		0.034	0.038	Experience, experience-squared, hours
		1994		0.050	0.074	Experience, experience-squared, hours
		1996		0.052	0.056	Experience, experience-squared, hours
		1998		0.051	0.077	Experience, experience-squared, hours
		2000		0.068	0.076	Experience, experience-squared, hours
	Clark (2003)	1995	0.062			Experience, experience-squared, gender, hours, state, tenure, tenure-squared
		1996	0.066			Experience, experience-squared, gender, hours, state, tenure, tenure-squared
		1998	0.118			Experience, experience-squared, gender, hours, state, tenure, tenure-squared
Slovak Republic	Chase (1998)	1984		0.028	0.044	Experience, experience-squared
		1993		0.049	0.054	Experience, experience-squared
		1984		0.031	0.048	Experience, experience-squared, industry, region
		1993		0.048	0.050	Experience, experience-squared, industry, region
	Filer, Jurajda, Planovsky (1999)	1995		0.063		Age, age-squared, industry, region, firm size, and the employer's ownership type
		1996		0.074		Age, age-squared, industry, region, firm size, and the employer's ownership type
		1997		0.081		Age, age-squared, industry, region, firm size, and the employer's ownership type

Source: As indicated.

Appendix Table 2. Descriptive Statistics By Specification: Pooled Samples

	Bulgaria				Czech Republic				Hungary				Latvia			
	Basic		Richer		Basic		Richer		Basic		Richer		Basic		Richer	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
Variables																
Ln(earnings)	8.72	0.76	8.72	0.75	8.73	0.46	8.74	0.46	9.97	0.57	9.88	0.56	4.16	0.64	4.20	0.61
Years of education	12.41	3.13	12.43	3.12	12.86	2.66	12.85	2.63	11.07	3.01	11.10	2.98	12.91	2.79	13.02	2.77
Experience	22.12	10.86	22.08	10.81	21.53	11.64	21.54	11.51	23.61	13.22	24.50	13.40	21.32	11.44	20.78	10.97
Male	0.50	0.50	0.50	0.50	0.49	0.50	0.49	0.50	0.47	0.50	0.46	0.50	0.46	0.50	0.46	0.50
Urban			0.78	0.41			0.65	0.48			0.58	0.49			0.68	0.47
Married			0.78	0.42			0.66	0.47			0.63	0.48			0.62	0.48
Occupation																
Plant and machine operators, etc.			0.45	0.50			0.58	0.49			0.57	0.50			0.52	0.50
Technicians and associate professionals			0.17	0.38			0.19	0.39			0.11	0.31			0.17	0.37
Professionals			0.22	0.41			0.11	0.31			0.10	0.30			0.17	0.38
Legislators			0.04	0.19			0.05	0.21			0.07	0.25			0.04	0.19
Public Sector			0.78	0.41			0.41	0.49			0.76	0.43			0.54	0.50
Trade Union membership			0.39	0.49			0.31	0.46			0.36	0.48			0.25	0.44
Full-time employee			0.89	0.31			0.91	0.29			0.69	0.46			0.87	0.34
Household size			3.43	1.33			3.04	1.22			3.45	1.63			3.13	1.29
Spouse: full-time employee			0.55	0.50			0.56	0.50			0.44	0.50			0.45	0.50
Number of observations	2,236		2,075		3,895		3,541		12,732		6,984		3,697		3,325	
Number of years	6		6		8		8		16		12		7		7	

Appendix Table 2 (continued)

	Poland				Russia				Slovak Republic				Slovenia			
	Basic		Richer		Basic		Richer		Basic		Richer		Basic		Richer	
	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev	Mean	Std Dev
Variables																
Ln(earnings)	6.12	0.53	6.15	0.51	5.38	1.10	5.72	0.88	8.54	0.47	8.59	0.48	10.77	0.70	10.86	0.66
Years of education	11.69	2.74	11.70	2.73	12.65	3.09	12.24	2.86	12.45	2.58	12.64	2.64	10.84	3.13	11.02	3.03
Experience	21.41	10.06	21.23	10.01	20.54	11.69	20.82	11.36	20.81	12.91	22.74	12.60	24.64	13.19	23.33	12.48
Male	0.50	0.50	0.50	0.50	0.48	0.50	0.49	0.50	0.48	0.50	0.46	0.50	0.49	0.50	0.49	0.50
Urban			0.73	0.44			0.76	0.42			0.37	0.48			0.60	0.49
Married			0.75	0.43			0.70	0.46			0.69	0.46			0.76	0.42
Occupation																
Plant and machine operators, etc.			0.56	0.50			0.52	0.50			0.50	0.50			0.56	0.50
Technicians and associate professionals			0.17	0.38			0.13	0.34			0.25	0.43			0.18	0.38
Professionals			0.12	0.33			0.21	0.41			0.08	0.28			0.11	0.31
Legislators			0.05	0.22			0.04	0.20			0.05	0.22			0.05	0.22
Public Sector			0.66	0.47			0.90	0.30			0.58	0.49			0.79	0.40
Trade Union membership			0.28	0.45			0.67	0.47			0.27	0.45			0.56	0.50
Full-time employee			0.92	0.27			0.89	0.31			0.79	0.41			0.79	0.41
Household size			3.77	1.63			3.35	1.23			3.18	1.26			3.59	1.30
Spouse: full-time employee			0.46	0.50			0.49	0.50			0.49	0.50			0.54	0.50
Number of observations	5,002		3,337		9,940		6,521		2,895		1,181		5,276		4,324	
Number of years	11		8		11		11		4		2		9		9	

Data Appendix

Our data are drawn from the *International Social Survey Programme* (ISSP), an ongoing annual program involving collaborative, international survey data collection efforts. Conducted annually since 1985, the surveys initially covered a narrow sample of industrial countries but, over time, have expanded to include more transition economies as well as developing countries. The ISSP currently consists of 32 national cross-sectional surveys. As described in the main text, the ISSP surveys are based on a common sampling and methodological framework and are thus comparable both between and within countries over time. The sample stratification generates nationally representative samples and provides individual-level information on demographic and socio-economic characteristics and personal views on selected social topics (such as social networks, work orientations, and gender roles) for at least 1,000 respondents per country per year.

Country and period coverage. The database used in this analysis consists of over 70 individual-level survey datasets for 8 transition economies (Bulgaria, Czech Republic, Hungary, Latvia, Poland, Russia, Slovak Republic, and Slovenia) over the 1986 to 2002 period, although Hungary alone has enough relevant information before the early 1990s. In particular we cover the following periods for each country: Bulgaria (1992-2002), Czech Republic (1993-2002), Hungary (1986-2002), Latvia (1996-2002), Poland (1991-2002), Russia (1991-2002), Slovakia (1995, 1998), and Slovenia (1991-1992).

Sample restrictions. We drop the following country-year surveys from the estimation sample. First, we drop ISSP data based on simultaneous surveys for Bulgaria (1996 and 1997; 1998 and 1999) and Slovenia (1993 and 1994; 1998 and 1999). In these simultaneous surveys, the modules differ but the core socio-economic background variables on which our analysis is based, is the same. We therefore exclude the duplicate surveys from the analysis. Second, we drop Russia in 1998 because the earnings data are based on midpoints, thus yielding very limited variability. Third, we drop the Czech Republic in 2000 because it only includes very few observations on monthly income. Fourth, we drop Poland in 1987 because there are no observations on years of education. After excluding self-employed workers, retirees, and students, our net sample size is about 500 individuals (age 18 to 65) per country per year. Because of the relatively small sample size, we are constrained from making gender-specific estimations.

Education and earnings variables. The crucial variables in our analysis are earnings and years of education. We discuss here possible differences in definition, scope, and comprehensiveness across countries.

With respect to schooling, respondents are typically asked the total number of years they attended school, although the specific phrasing varies somewhat from country to country. In the Czech Republic, for example, they are asked, "How long did you go to school (include also incomplete or part-time studies)?" In Russia, "How many years altogether did you study? (including vocational training)." The survey questions do not explicitly instruct respondents to include courses or years of study that were not completed; the one exception is the Czech Republic.

As for earnings, respondents are typically asked about their net, rather than gross, monthly earnings, except in Russia where the wording of the survey question is somewhat ambiguous. In Russia, respondent are asked about their "monthly wage together with all the bonuses, compensations and other payments." In other countries, however, the survey question explicitly asks about net monthly earnings/income. In particular, they are asked about "personal monthly net income" (Bulgaria), "total average net income per month" (Czech Republic), "net earnings per month" (Hungary), "monthly net income" (Latvia), "average monthly income/earnings from your job or business after taxes" (Poland), and "last regular monthly income after taxes and social insurances from all sources and social insurances from all sources" (Slovenia). From the last example (Slovenia), it is also clear that the comprehensiveness of earnings/income varies across countries. In the Czech Republic, Russia, and Slovenia, respondents are explicitly asked about monthly income/earnings from all sources. In fact, in the Czech Republic and Slovenia, they are asked to include unemployment benefits, maternal benefits, and social insurance. In contrast, in Bulgaria, Hungary, Latvia, and Poland, respondents are simply asked what their monthly net incomes/earnings are.

Thus, these two key characteristics of the earnings data—(a) whether they are before of after taxes; and (b) whether they refer to the main job alone, earnings from all earnings-generating activities, or income from all sources including social transfers—suggest that some part of the differentials in estimated returns to schooling may be due to differences in progressive taxation or in the comprehensiveness of income/earnings data. Within countries, transfers may also distort the estimated returns to schooling. Where there have been substantial reforms in the tax and transfer system, estimating the returns to schooling in a consistent manner is particularly difficult. However, the restrictions we impose on our sample and the adjustments we make to our data mitigate somewhat the cross-country differences in the definition of earnings. For example, we restrict the sample to the wage-employed, thus making an effort to ensure that our comparisons within and across countries are of wages and earnings, not of entrepreneurial income or income from self-employment. Restricting the sample to those currently employed also effectively excludes those currently receiving unemployment benefits and allowances, or those on maternity leave receiving maternity benefits, although we are unable to identify those who are registered unemployed (and thus may be receiving benefits), while working informally. In addition, an auxiliary set of regressions uses hourly earnings rather than monthly earnings; this helps facilitate the comparison of those with single jobs with those performing multiple jobs.

Sector of employment. A section of the main text compares the levels of and evolution of returns to education in the public sector with those in the private sector. The absence of a clear definition of sector of employment often complicates the analysis of this dimension of transition. In our analysis, we classify as public sector employees those who declared themselves working for the government sector, the public sector, and for a publicly owned firm, including nationalized and state owned enterprises. The treatment of partial ownership in ISSP, however, is not clear. Hungary explicitly and consistently includes companies with mixed ownership under "publicly-owed firms." In contrast, the Czech Republic, at least in the mid-1990s, considered firms with mixed ownership as private.

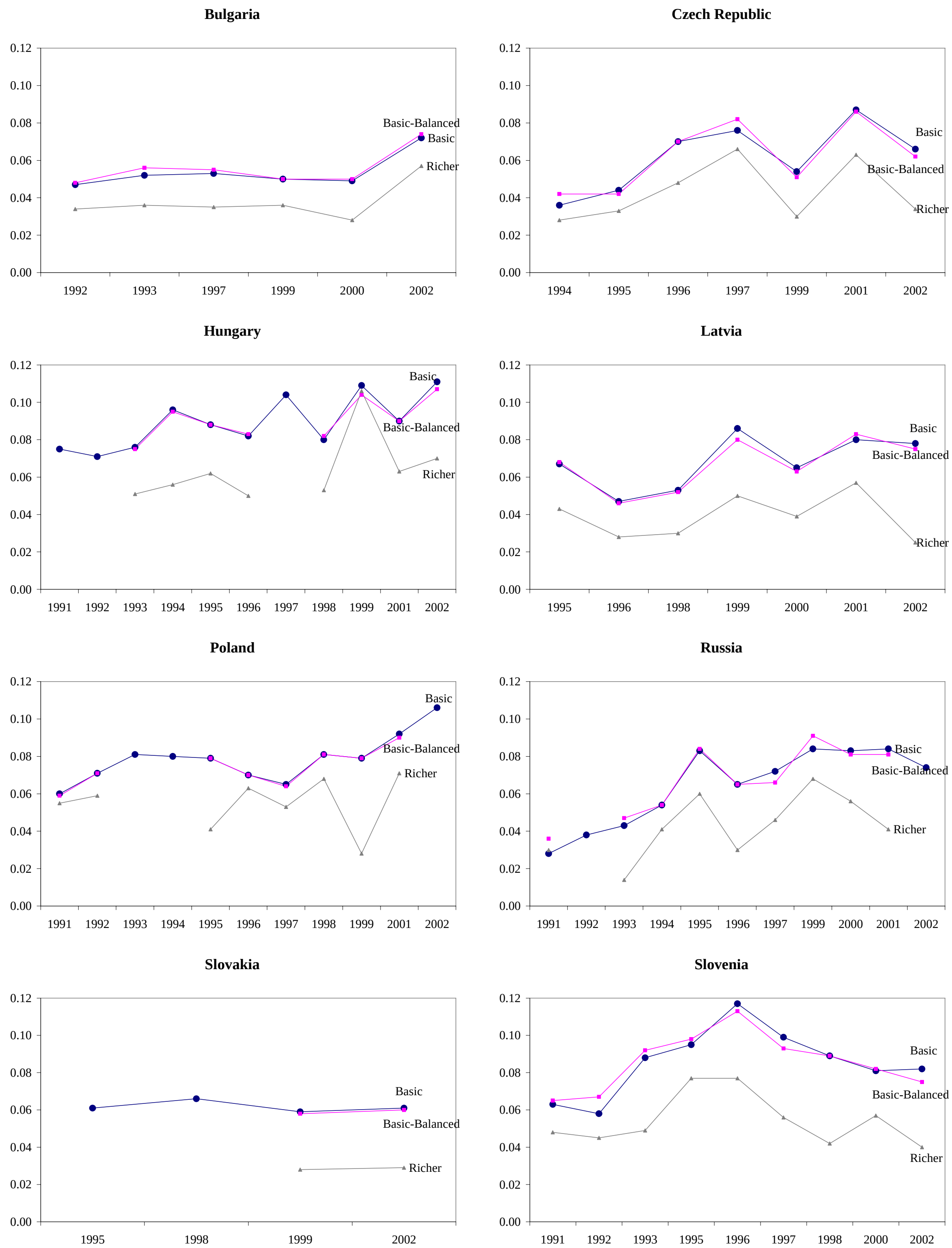
Participation rates. Cross-country differences in returns to schooling may be driven in part by differences in labor force participation (LFP) rates. Because ISSP data do not allow for a careful inspection of ILO-consistent measures of labor force participation rates, we rely on data from the Key Indicators of the Labor Market (KILM) database (Appendix Table 3). There are breaks in the series through the early transition period, so the LFP rates from the mid-1990s onwards are typically measured more consistently. Bulgaria and Russia, however, have breaks throughout most of the entire period and their LFP rates should be interpreted with caution. Nonetheless, the data suggest that LFP rates have remained more or less constant from the mid-1990s through 2002. On average, LFP participation rates are 66-67 percent, with some variation across countries. The Czech Republic is at the higher end of this distribution (over 70 percent) while Hungary is at the lower end (about 59 percent). On average, men have higher participation rates, over 70 percent on average since the mid-1990s compared to about 60 percent among women.

Appendix Table 3. Labor Force Participation Rates

	1990	1991	1992	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002
All													
Bulgaria	75	...	68.8	...	...	74.2	...	...	...	...	58.9	62.4	...
Czech Republic	...	...	...	72.1	72.4	72.3	72.1	72.1	72.2	72.2	71.6	71.1	70.9
Hungary	65.7	...	64.4	62	60	58.9	58.5	57.8	58.4	59.9	59.9	59.6	59.7
Latvia	79.4	...	...	...	...	79.2	70.5	70.2	69.5	68.7	67.6	67.7	...
Poland	72.5	...	69.4	68.8	68.4	67.4	66.9	66.4	66.1	65.9	65.8	65.7	64.8
Russia	76.5	...	72.6	...	...	69.4	67.4	...	66.7	68.9	...	...	...
Slovakia	78.2	76	...	71.9	69.3	69.3	69.8	69.4	69.3	69.5	69.9	70.5	69.9
Slovenia	70.7	68.6	...	60.7	66.3	67.8	66.8	67.9	69.1	68	...	67.8	...
Men													
Bulgaria	77.7	...	71.3	...	...	77	...	...	...	...	63.3	66.3	...
Czech Republic	...	...	...	80.3	80.4	80.6	80.7	80.5	80.3	80.2	79.4	79	78.9
Hungary	74.5	...	71.9	69.4	67.8	67.9	67.4	66.6	66.3	67.8	67.5	67.2	67.1
Latvia	83.6	...	...	...	...	83.2	77.6	76.5	76.6	75.3	72.5	72.9	...
Poland	80.1	...	76.4	75.7	75	73.9	73.5	73.2	72.8	72.3	71.7	71.5	70.8
Russia	81.6	...	76.8	...	...	76.2	74.3	...	...	74.2	...	...	...
Slovakia	82.5	81.5	...	78.3	77.6	77.3	77.1	76.9	77.2	76.9	76.8	77.4	76.7
Slovenia	76.8	73.7	...	64.7	71.6	73.4	71.5	72.3	73.3	72.2	...	72.2	...
Women													
Bulgaria	72.2	...	66.2	...	...	71.6	...	...	...	...	54.6	58.7	...
Czech Republic	...	...	...	63.9	64.4	64.1	63.6	63.7	64	64.1	63.7	63.2	62.8
Hungary	57.3	...	57.3	54.9	52.7	50.3	49.9	49.3	50.8	52.3	52.6	52.4	52.7
Latvia	75.6	...	...	...	...	75.4	63.9	64.5	63	62.6	63	62.8	...
Poland	65.1	...	62.6	62.1	62.1	61	60.5	59.9	59.7	59.8	59.9	59.9	58.9
Russia	71.7	...	68.6	...	...	63	61	...	...	63.9	...	...	...
Slovakia	74.1	70.7	...	65.7	61.3	61.4	62.5	62	61.7	62.3	63.2	63.8	63.2
Slovenia	64.7	63.4	...	56.3	61	62.6	61.9	63.2	64.8	63.3	...	62.8	...

Source: KILM database.
Includes the working-age (15-64) population only.

Figure 1. Returns to Education By Specification



Note: Coefficients of the years of schooling variable in earning regressions. Dependent variables are monthly earnings. Robust standard errors in parentheses. Specifications include the following regressors. Basic: potential experience (linear and squared), dummies for male. Richer: previous regressors plus dummies for living in urban areas and married, controls for current job (dummies for occupation, public employee, working full-time, member of a trade union), controls for current family (number of members, dummy for spouse working full-time).